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Research Article

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Michaela Kreyenfeld

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Michaela Kreyenfeld [1]

Abstract

This paper investigates the role of women's education in the transition to the second child using data from the 1997 German micro-census. We begin our analysis with a simple model, which shows a positive effect of woman's education on the transition rate to the second child for West German women. We argue that this effect is most likely confounded by various factors. Firstly, we assume that there is a *time-squeeze* effect, which increases the transition rate to the second child for more highly educated women. Secondly, titled as the *partner hypothesis*, we argue that more highly educated women often live with more highly educated partners who have the earning potential to afford a large family. Thirdly, titled as the *selection hypothesis*, we argue that the positive effect of women's education can be attributed to a selection effect, i.e. family-oriented college graduates are more likely to select themselves into the group of women at risk of second birth. The empirical investigations particularly support the second and third hypotheses. After controlling for the partner's characteristics and including unobserved heterogeneity factors, the positive effect of female education becomes strongly negative.

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1. Introduction

In the economic theory of fertility, considerable attention has been paid to the role of women's employment in fertility (e.g., Becker 1993, Oppenheimer 1994). A major assumption that is made –frequently labeled as the "male chauvinist approach"– is that the woman in the household gives up work on the labor market in order to take care of the children. This might be the result of intra-family bargaining, which assigns the childcare responsibilities to the person with the lowest market wage (e.g., Ott 1992), or it might just as well be an indicator of the persistence of traditional gender roles. However that may be, assuming that the care for children is considered a woman's responsibility and assuming further that rearing children and employment are incompatible, high female wages should increase the opportunity costs of childrearing and suppress fertility. Assuming further that a woman's educational attainment is a valid indicator for her work orientation, her wage and her labor market chances, one would expect a **negative** correlation between female education and fertility.

In recent years, however, an increasing number of studies have reported a **positive** effect of women's education on the transition rate to higher order births. These studies primarily use data from Scandinavian countries (B. Hoem 1996, Hoem and Hoem 1989, Kravdal 1992, Oláh 1996), but similar findings have also been reported for western European countries such as Austria and West Germany (e.g., Hoem et al. 2001). In particular, for the West German context, there are the studies by Huinink (1989, 1995, 2001) which address the role of women's educational attainment for the transition to the second child. Using data from the German Life History Survey, Huinink (1989, 1995) finds that women with an *Abitur* (German high school degree) encounter a higher second birth risk than women with only a *Realschulabschluss* (intermediate school degree). In a more recent publication, Huinink (2001) uses data from the Family and Fertility Survey and cross-tabulates the completed family size by women's educational level for various European countries. For West Germany, he finds a high percentage of childless women among the college graduates, but also a high percentage of women with two or more children.

A straightforward way to explain such a pattern could be that women's education positively affects fertility through an **income effect**. More highly educated women earn higher wages, they actively contribute to the household income and should therefore be better able to support a larger family (Kravdal 1992: 468, Macunovich 1996, Rindfuss et al. 1996: 280). If this interpretation were correct, it would indicate a path-breaking social change, involving a role change for women from caregivers to household providers. Furthermore, it could mark a turning point in demographic development. While it has been argued that woman's increasing education and economic

independence has contributed to the past downward trend in fertility (see e.g., Brewster and Rindfuss 2000: 271), such an interpretation suggests an upward trend for the future.

The pivotal assumption, which the "income effect hypothesis" implicitly relies on, is that childrearing and employment can be made compatible and that women are able to return to work after childbirth. Women must be able to bring an employment career in synch with childrearing. This assumption may be plausible for Scandinavian countries, where full-time day care is readily available, but it is not very convincing for West Germany (Note 1). The West German institutional framework is frequently cited as the prototype of the "male breadwinner model" (e.g. Esping-Andersen 1999). Public daycare is scarce and there are only limited chances to arrange daycare by relying on private modes of care. Against this background, it is puzzling that an "income effect" of women's education would foster childbearing.

The primary goal of this paper is to investigate alternative hypotheses that might explain a positive effect of women's education on the transition rate to the second child. Woman's education is measured in this paper by a set of three binary variables, which indicate whether the respondent has a college degree (*Universitätsabschluss, Fachhochschulabschluss),* a vocational degree *(Ausbildungsabschluss, Fachschulabschluss)* or no higher educational degree whatsoever. The remainder of this paper is structured as follows. In Part 2, we discuss the institutional constraints for mothers participating in the labor market in Germany. We furthermore discuss a "time-squeeze", a "partner effect" and a "selection hypothesis" as possible explanations for the unexpected positive effect of women's education on second birth risks. Part 3 presents the data set and describes the empirical procedure. Part 4 comprises the empirical analysis and Part 5 consists of our concluding remarks.

2. Theoretical Considerations

In cross-national comparisons, the German welfare state is often characterized as fostering the traditional "male breadwinner model" (e.g., Gauthier 1996: 155, Gornick et al. 1998, Sundström 1999). There are various institutional regulations that have made researchers arrive at such a conclusion. For example, the system of income splitting *(Ehegattensplitting)* allows married couples to file their taxes jointly, i.e. the man and the woman's income are added together, divided by two and then taxed as individual incomes. Since there is also a progressive tax schedule, the tax relief is particularly high for couples in which one of the partners is permanently not employed or only employed part-time. The German health care and pension systems contain similar incentive structures. Married housewives are automatically insured by the health insurance of their partners *(Familienmitversicherung)* and entitled to a widow's pension.

In theory, this system is neutral towards gender roles and it does not necessarily force the woman to withdraw from the labor market after childbirth. In practice, however, it is primarily the woman who gives up employment to raise children. For example, parental leave is almost exclusively (in 98 percent of all cases) taken by women (Engelbrech 1997). Furthermore, due to a lack of children's day care, women's ability to combine childrearing and employment are limited. The provision rate of public day care for children ages 0-3 is a mere three percent, while for children ages 7-10 it is only six percent. For pre-school children (ages 4-6), there is a broad coverage of part-time care, but the provision rate of places in full-time care is only 16 percent (Statistisches Bundesamt 2001). The opportunities to find private daycare arrangements *(Tagespflege)* are scarce, too. High market barriers of entry, wooly quality regulations and high tariff wages for childcare employees are presumably major obstacles that have prevented a functioning private market for care from evolving (for details, see e.g., Kreyenfeld and Hank 2000).

Since childcare opportunities are scarce, women are often confined to –at least temporarily– withdrawing from the labor market when they have children. Table 1 exemplifies this issue by displaying the employment rates of West German mothers by the age of the first child (Note 2). When the first child is between 3 and 6 years old, only around 10 percent of the West German women are employed full-time (for a longitudinal analysis, see e.g., Drobnič 2000, Kurz 1998). There is some variation in the full-employment rates by the educational level of the woman. College educated women are more likely to work full-time than non-college educated women. Nevertheless, the full-time employment rate for college educated women with children in pre-school age (4-6 years of age) is still only roughly 20 percent.

Table 1: *Employment rate of mothers in West Germany in 1997*

	Age of first child		
	0-3	**4-6**	**7-10**
No degree*)			
Employed full-time	5%	9%	8%
Employed part-time	9%	15%	21%
Not employed	86%	76%	71%
Sample size	*1,131*	*1,155*	*1,349*
Vocational degree*)			
Employed full-time	9%	10%	11%
Employed part-time	17%	32%	41%
Not employed	75%	58%	48%
Sample size	*3,615*	*3,243*	*3,439*
College degree*)			
Employed full-time	14%	18%	21%
Employed part-time	25%	33%	34%
Not employed	61%	49%	45%
Sample size	*536*	*384*	*288*
All			
Employed full-time	8%	10%	10%
Employed part-time	15%	27%	35%
Not employed	73%	61%	53%
In education	3%	2%	2%
Sample size	*5,452*	*4,884*	*5,173*

Notes: (1) Population: women of the birth cohorts 1961-1980 whose youngest child is age 0-10 in 1997
(2) *) Respondents who are in education are omitted
(3) Employed full-time is defined as working 35 or more hours per week. Employed part-time is defined as working 1-35 hours per week.

Source: *Mikrozensus* 1997

The restricted opportunities to find out-of-home care and the system of income splitting are presumably the glaring characteristics of (West) German family policies. Huinink (2001) argues that this specific institutional framework has contributed to a "polarized" fertility behavior in West Germany. Since childrearing and employment are barely compatible, women are often confined to choose between motherhood and an employment career as two exclusive alternatives. If they decide for the "family track"

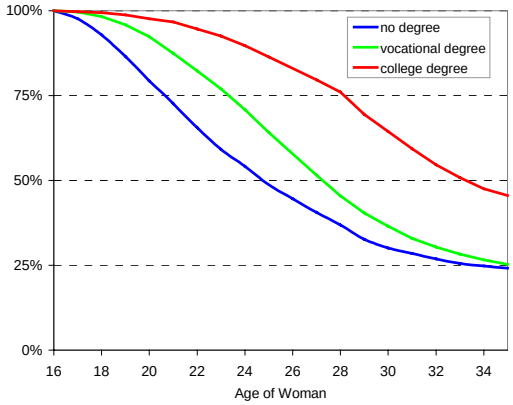
by having a first child, they simultaneously opt out of the labor market. If they decide for the "career track", they will ultimately remain childless. Since more highly educated women generally encounter better career opportunities, they should be more likely to belong to the latter category.

Figure 1 gives support for the view on West German fertility. It shows the survival curves for the transition to the first child by final educational attainment for the birth cohort 1961-1963 (which will also be used for the subsequent analysis). While 25 percent of the women without a college degree are still childless at age 35, this applies to almost 50 percent among the college graduates.

The intriguing question is how the West German institutional framework shapes second birth patterns. As noted above, studies from other countries report a positive effect of women's college education on the transition rate to higher order births. One plausible explanation for such a pattern is an **income effect**. More highly educated women earn higher wages and they are therefore better able to afford a larger family. Another explanation for this pattern could relate to **work accelerated childbearing** *(*Ní Bhrolcháin 1986a, 1986b). It might simply be sensible for work-oriented women to space their births close together. This allows them to swiftly resume employment, which reduces childcare related employment interruptions, minimizes both forgone earnings and risks of a devaluation of human capital (Taniguchi 1999). While this argumentation is plausible for countries where childrearing and employment is compatible, it is not convincing against the background of the West German institutional framework.

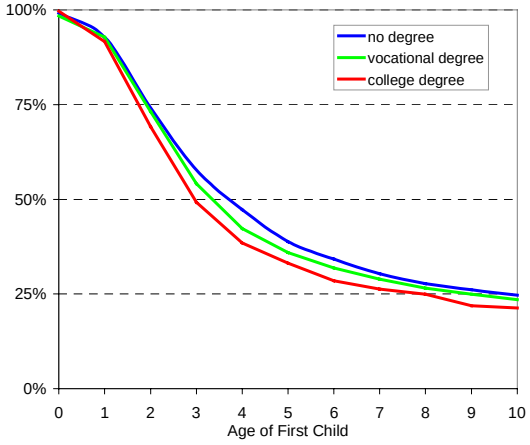
Figure 2 shows the survival curve to the second child by the woman's final educational level for West German women of the cohorts 1961-1963. In contrast to the first birth pattern, there are hardly any differences in the final progression ratio by the woman's educational level. Ten years after the birth of the first child, roughly 75 percent have another child. College graduates, however, have the second child slightly more rapidly than women of the other educational categories. In other words, similar to the studies from other countries, college graduates encounter a higher transition rate to the second child in West Germany.

Figure 1: *Transition to the first child by final educational attainment of the woman (survival curve)*



Population: German nationals of the birth cohort 1961-1963 who are living in West Germany
Source: *Mikrozensus* 1997 (own estimates)

Figure 2: *Transition to the second child by final educational attainment of the woman (survival curve)*



Population: German nationals of the birth cohort 1961-1963 who are living in West Germany and who have at least one child
Source: *Mikrozensus* 1997 (own estimates)

In the following, we more closely investigate three aspects which might contribute to this positive effect of women's college education on second birth risks (Note 3):

1. College educated women are usually older when they have their first child than other women. It usually requires a longer period of time to receive a college degree than is the case for receiving a vocational training certificate. Since participation in the educational system is not easily compatible with raising children, college graduates generally postpone parenthood until they reach mature ages (see e.g., Blossfeld and Huinink 1991). For example, in our data set (see below), the average age at first birth for women with a vocational certificate is 25.7, while for college graduates it is 28.1. Having a first child later in one's life involves having less time at one's disposal before reaching the biological limits of fertility. Such a **time-squeeze** could increase the transition rate to the second child.

2. The second aspect refers to "educational homogamy", which is relatively common in Germany. (For a detailed discussion, see Wirth 2000.) For example, 60 percent of all women in our sample live with a partner with the same educational level, whereas only 27 percent are "heterogamous couples" and 13 percent do not live with a partner at all (see Table 2). Educational homogamy can be very important if one investigates the role of women's education in fertility. In a "male breadwinner regime" like the (West) German one, the male's employment situation and income level should be crucial for fertility decisions. Couples would be likely to postpone parenthood until the man has established a secure position in the labor market, and they should only opt for a second or third child if they see the man's earnings as sufficiently high to support a larger family. Assuming that education is a valid indicator for the "earning potential" of individuals, women who have college-educated partners should encounter higher transition rates to second or third births. If there is a high degree of educational homogamy and one fails to control for the partner's characteristics, one might get biased results. The effect of a woman's college education might then be confounded with a **partner effect**.

3. Finally, women who are at risk of second birth are a select group of individuals because they must already have one child. Therefore, they have manifested some preference for children and committed themselves to a child-oriented "life plan". We argue that this "selection effect" matters particularly for the role of education in the risk of having a second child. Given the unfavorable economic constraints for combining childrearing and employment in West Germany, women with a college degree face high opportunity costs if they decide to have a first child. College graduates who nevertheless opt for parenthood probably have other characteristics (such as an extraordinarily high preference for children) that have contributed to

such a decision. In other words, the positive effect of women's college education might be far from "real", but merely attributable to **self-selection**.

In the following, we investigate these aspects. For this analysis, we are using data from the scientific-use file of the 1997 German micro-census. Let us first describe the data set.

3. Data and Method

3.1 Selection of the Sample

The German micro-census (hereafter referred to as *Mikrozensus*) is a one-percent sample of the population residing in Germany and it covers standard socio-demographic characteristics such as age, nationality, region of residence, educational attainment, etc. (For details, see e.g., Emmerling and Riede 1997, Schimpl-Neimanns 1998.) The scientific-use file of the *Mikrozensus* is a 70 percent sample of the original *Mikrozensus* and therefore a 0.7 percent sample of the population residing in Germany. The major advantage of the *Mikrozensus* is that it is a highly reliable and representative data set, and that it provides a large number of cases. The entire *Mikrozensus* for the year 1997 contains roughly 500,000 respondents.

Before we discuss the selection of the data set, it is important to draw attention to some shortcomings of the *Mikrozensus*. Its major drawback in regards to the analysis of demographic events is that it does not provide the "fertility history" of the respondents. The date of birth and the parity of the children can only be inferred from the number and age of the children who live in the household at the time of the interview. Fortunately, comparisons with other data sets show that this is only a minor problem for the birth cohorts born after 1960. However, for older cohorts, the percentage of children who have already moved out of the parental home increases rapidly (Kreyenfeld and Huinink 2001) (Note 4). For this reason, we restrict our analysis to the birth cohorts born in 1961 or later. We furthermore omit from our analysis the cohorts born later than 1963 because we must concentrate on respondents who have a first child. All in all, this leaves us with the birth **cohorts 1961-1963**, who were between the ages of 34 and 36 at the time of the interview (in 1997). This obviously does not allow us to investigate any cohort effects, but it provides us with a very homogenous sample.

We furthermore restrict the analysis to respondents who live in **West Germany**. To include East Germany would have complicated the analysis, since we would have to take into consideration the change of regime in 1990 (see Kreyenfeld 2001). We furthermore exclude foreign nationals from the analysis. Other, more minor restrictions

are the following: We restrict the analysis to respondents in private households, i.e. we omit respondents who primarily live in institutions (such as mental asylums). We also exclude all cases where a birth occurred before age 17. Furthermore, we exclude cases with missing information on the educational attainment of the woman or her partner. Finally, we omit women for whom a first and a second birth were recorded in the same calendar year. We do not have any information on the month of birth at our disposal, so we must assume that such children are twins (Note 5). We omit them from the multivariate analysis since it is not possible to calculate a positive duration for them. The remaining sample for the analysis of second births comprises 6,026 women who gave birth to 4,035 second children (see Table 2).

3.2 Description of the Variables

The key independent variables for our multivariate analysis are the educational levels of the female respondent and her partner. We use the highest educational level received at the time of interview. One could argue that it would have been more advantageous to use education as a time-variant covariate (J. Hoem 1996, Hoem et al. 2001: 252). However, for the analysis of second birth it is relatively trouble-free to use the highest educational level, since most respondents have completed their studies before the first child is born, respectively before being at risk of second birth. As said before, we make a distinction between respondents with a college degree *(Universitätsabschluss, Fachhochschulabschluss),* a vocational degree *(Ausbildungsabschluss, Fachschulabschluss)* and respondents with neither a college nor a vocational degree. We labeled the latter category "no degree". It is worth noting that we do not consider people with a primary or secondary school degree as a separate group. This means that we might have classified some respondents into the category "no degree" even if they received an "Abitur". The main reason we proceeded this way is that in the German labor market, vocational training certificates as well as college degrees play more significant roles for the allocation of workers to jobs than secondary school degrees do. Therefore, they better indicate career opportunities, earning potential, unemployment risks, etc. (see e.g., Müller and Shavit 1998).

Table 2 displays the distribution of the sample. It also displays the distribution of the women who are at risk of first birth, which we will draw on when we investigate the "selection hypothesis". As can be seen from the table, roughly 10 percent of the respondents have a college degree. This is slightly below the average ratio of women with a college degree in the total population, reflecting the lower probability of college educated women to have a first child. About 17 percent of the women in the sample live with college-educated partners, 59 percent with partners with a vocational degree and

only 10 percent with partners without any degree. The remaining 13 percent do not have partners at the time of interview (Note 6). For some parts of the analysis, we combine the woman's education with her partner's. This is why the table also reports the combination of the woman's and her partner's education.

Table 2: *Composition of the sample*

	First birth	Second birth
Mean age at first birth		
No degree	--	23.69
Vocational degree	--	25.69
College degree	--	28.06
All	--	25.55
Woman's education		
No degree	18%	19%
Vocational degree	68%	71%
College degree	14%	10%
Partner's education		
No partner	15%	14%
No degree	8%	9%
Vocational degree	53%	61%
College degree	16%	17%
Woman & partner's education combined		
No partner	23%	14%
Educational homogamy	54%	60%
Woman with higher education	8%	6%
Partner with higher education	24%	20%
Number of cases		
Sample size	8,530	6,026
Number of births	6,026	4,035

Notes: Population: Women of the cohorts 1961-1963. Cases with missing information on the woman's or her partner's education level were omitted.
Source: *Mikrozensus* 1997

3.3 Analytical Procedure

As we have said before, we hypothesize that the model used for the empirical analysis of the role of women's education in second birth risks is easily subject to misspecification. We briefly discussed three aspects which could confound this relationship: (1) a time-squeeze, (2) a partner effect and (3) self-selection. In the following, we investigate, one after the other, the three issues sketched above. We estimate several event-history models. In all models, the duration since the birth of the first child is used as the baseline hazard.

When $h(t)$ is the hazard of occurrence at time t, β_1 the baseline hazard, X the (time-constant) covariates, and β_2 a vector of corresponding parameters, the equation to be estimated has the following general mathematical form

$$ln\ h(t)=\beta_1(t)+\beta_2 X \qquad (1)$$

The baseline log-hazard is a piecewise-linear spline (also known as a generalized Gompertz function). The parameters are estimated using the software aML (Version 1.04). This software has a variety of virtues which will become apparent in the subsequent analysis. First, aML allows inserting continuous covariates (in our case the woman's age at first birth) as a piecewise linear function. A linear spline is a flexible form of representing the effect of a continuous independent variable. Such a variable is basically "cut" into several segments. For each segment, the log-hazard of the independent variable is assumed to be linear. However, across segments, the parameters (which appear as slopes of the spline function) can be different. (For details, see Lillard and Panis 2000: 46.) In order to interpret the slope coefficients, it is easiest to visualize them in a graph (see below). Apart from using a log baseline and one or more covariates represented as a linear spline, aML also allows us to control for unobserved heterogeneity. This aspect will be of particular importance when investigating the "selection hypothesis".

4. Results

4.1 Plain Model

In a first step, we estimate a "plain model" where we solely control for the educational level of the respondent. Table 3 reports the results from this model. Having a college degree increases the relative risk of having a second child by 10 percent compared to respondents with only a vocational training certificate. Compared to women without

any degree, it increases the relative second birth risks by 17 percent. This is basically the results we would have expected based on the survival curves (see Figure 2) and based on previous studies on the role of women's educational attainment on second birth risks in West Germany (see Huinink 1989, 1995) (Note 7).

Table 3: *Event-history model of the transition to the second child, Model (0): plain model*

	β	exp(β)	t	
Baseline (ls)				
Intercept	-4.32	0.01	-40.18	***
0-3 years	1.22	3.38	27.97	***
3-5 years	-0.72	0.49	-20.55	***
5+ years	-0.15	0.86	-8.57	***
Woman's education				
No degree	-0.06	0.94	-1.49	
Vocational degree	0	1		
College degree	0.10	1.10	1.80	*
Log Likelihood	-8,009			

Notes: (1) Method: event-history model (generalized Gompertz)
(2) time variable: age of first child
(3)***: $p \leq 0.01$ **: $0.01 \leq p \leq 0.05$ *: $0.05 \leq p \leq 0.10$
Source: Mikrozensus 1997 (own estimates)

4.2 Time-Squeeze Hypothesis

We now turn to our first hypothesis. Since college educated women are often older at first birth, they have less time left to have their second children before reaching the biological limits of fertility. This might induce them to more rapidly opt for the second child to reach their desired number of children. In order to investigate whether a late age at first birth accelerates second birth risks, we add the age at first birth to the model. In Model (1a), we insert the age at first birth as a categorical variable, in Model (1b) as a piecewise linear spline. The two specifications provide different kinds of advantages. The use of a categorical variable is, on the one hand, very simple, since it allows us to employ interaction effects relatively easily (see below). Using a (regressor) spline, on the other hand, is a more flexible type of modeling.

As can be seen from Figure 3, the two kinds of specifications provide fairly similar results concerning the role of a woman's age at first birth in second birth risks. Women who had their first child as teenagers have relatively low transition rates to the second child. If one ignores this group of women, the relationship between the age at first birth and the transition rate to the second child is basically negative. There is no significant difference between women who had their first child between the ages of 21 and 25 and those who were between the ages of 26 and 30 at this event. However, when the woman was older than age 30 at the birth of her first child, the transition rate to the second child declines rapidly.

Table 4: *Event-history model of the transition to the second child, Model (1a+b): the role of the age at first birth*

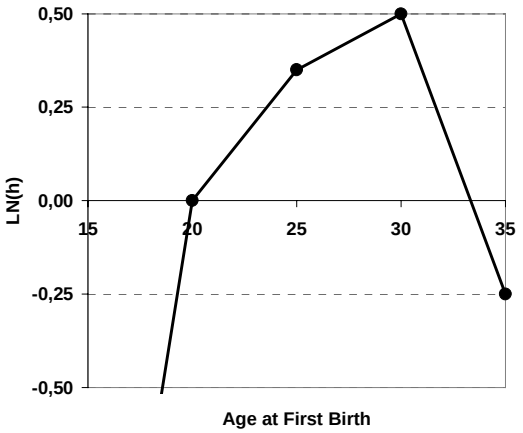
	Model (1a)				**Model (1b)**			
	β	exp(β)	t		β	exp(β)	t	
Baseline (ls)								
Intercept	-4.23		-38.67	***	-4.57		-34.14	***
0-3 years	1.21		27.71	***	1.20		27.49	***
3-5 years	-0.74		-20.92	***	-0.73		-20.63	***
5+ years	-0.15		-8.28	***	-0.13		-7.11	***
Woman's education								
No degree	-0.06	0.95	-1.35		-0.04	0.96	-0.95	
Vocational degree	0	1			0	1		
College degree	0.13	1.14	2.33	**	0.14	1.15	2.58	***
Age at first birth								
17-20	-0.10	0.90	-2.08	**	0.35 [ls]		0.98	
21-25	0	1			0.07 [ls]		3.23	***
26-30	-0.03	0.97	-0.86		0.03 [ls]		2.55	***
31-35	-0.41	0.66	-5.24	***	-0.15 [ls]		-8.98	***
Log Likelihood		-7,993				-7,949		

Notes: (1) Method: event-history model (generalized Gompertz)
(2) time variable: age of first child
(3) ***: $p \leq 0.01$ **: $0.01 \leq p \leq 0.05$ *: $0.05 \leq p \leq 0.10$
(4) ls= linear spline

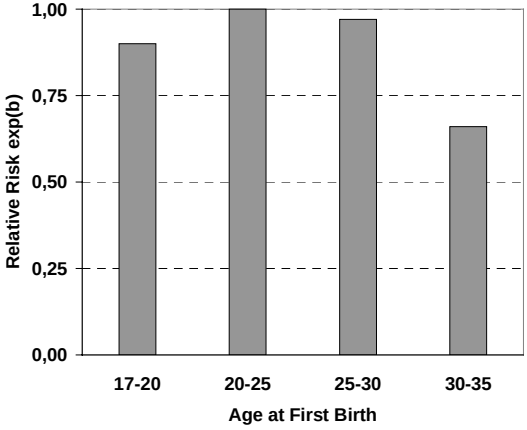
Source: *Mikrozensus* 1997 (own estimates)

***Figure 3**: Second birth risks by age at first birth*

***Panel 1**: Age at first birth as a linear spline*



***Panel 2**: Age at first birth as a categorical variable*



Notes: For the estimates, see Table 4.
Source: *Mikrozensus* 1997 (own estimates)

Our primary motivation for inserting the age at first birth was to investigate whether the effect of women's college education is confounded by the relatively late age at first birth for college educated women. After controlling for the age at first birth, we expected that the positive effect of women's college education would disappear. However, as can be seen from Table 4, after adding the age at first birth to the model, the positive effect of a woman's college education becomes highly significant and increases even slightly (compared to the model reported in Table 3).

However, by just inserting the age at first birth as a (multiplicative) covariate might conceal interaction effects between women's education and the age at first birth. A late age at first birth might generally have a depressing effect on the transition to the second child. This does not necessarily rule out that college graduates, who usually have a first child later in their lives, accelerate childbearing. In order to investigate the interaction between the age at first birth and the educational level, we proceed in a similar way as was suggested by previous studies (particularly by B. Hoem 1996).

Interaction between age at first birth and educational attainment

In our next step, we employ an interaction of the woman's educational attainment with the categorical variable for the age at first birth. For women who gave birth to a first child as teenagers, we do not allow the variable to vary by educational attainment. There are simply too few teenage births to college educated women to allow for such an interaction. The interaction effects of this model are displayed in Table 5.

There are two ways of reading the table. Firstly, one can compare the relative risks along the rows. In other words, one fixes the educational level in order to investigate whether a late age at first birth operates differently for the various educational categories. For all educational groups, there is a negative gradient of age at first birth on the transition to the second child. For the college graduates, however, this effect is less strong than for the other groups.

Secondly, one can read the table column-wise. One then compares the effect of education by the age at first birth. Given that the positive effect of women's educational attainment was primarily transmitted through a late age at first birth, it should disappear if one holds the age at first birth constant. As can be seen from Table 5, the effect of education disappears for the group of women who had their first child at the ages 21-25. For the other age categories, however, there is still a positive gradient. For the medium age group (first birth at the ages 26-30), relative first birth risks for college graduates increase by 26 percent compared to women with a vocational degree. In the oldest age category (first birth at the ages 31-35), relative second birth risk for college graduates increase by 39 percent (Note 8). In sum, the general picture remains basically

unchanged; women's education positively affects second birth risks, irrespective of the age at first birth.

Table 5: *Event-history model of the transition to the second child, Model (1c): Interaction between age at first birth and education*

	Age at first birth		
	21-25	**26-30**	**31-35**
	exp(β)	exp(β)	exp(β)
Woman's education			
No degree	0.94	0.85	0.45
Vocational degree	1	0.95	0.64
College degree	0.93	1.20	0.89

Notes: For the full model (without interaction effects) see Model 1a.

From our analysis so far, one could conclude that the relatively late age at first birth does not contribute to the positive effect of women's college education on second birth risks in West Germany. Neither do our investigations reveal any substantial interaction effects between women's educational attainment and the age at first birth. However, our "standard interaction procedure" might not fully take into account the connection between age at first birth and educational attainment in a satisfying manner. The main argument in this context is that the distribution of the age at first birth differs strongly by educational levels. Related to this, a late age at childbirth is defined in a completely different manner for college graduates than for others. "What is completely normal childbearing behavior for one educational group is quite unusual in another" (B. Hoem 1996: 337). In order to take into account that the age at first birth has a different meaning for the various educational categories, one should take into account the **distribution** of the ages at first birth by educational level.

In our next step, we use the mean age at first birth by educational level from Table 2 in order to construct an indicator variable for whether a woman's age at first birth is "below the group average" or "above the group average" (Note 9). We then employ an interaction of this variable with the educational level of the respondent. The results are displayed in Table 6. For the group of women who had the first child below the average of her educational level, there is indeed no effect of education. However, for the other group there is still a strong and significantly positive effect of women's educational attainment on second birth risks. In other words, even after taking into account the "relative age at first birth", the positive effect of women's college education remains. For Sweden and Austria, such a procedure revealed an accelerating effect of

the age at first birth on third birth risks which confounded the effect of women's education (B. Hoem 1996, Hoem et al. 2001). However, for West Germany we do not find evidence for such an effect.

Table 6: *Event-history model of the transition to the second child, Model (1d): Interaction between relative age at first birth and education*

	Age at first birth	
	Below average	**Above average**
	exp(β)	Exp(β)
Woman's education		
No degree	0.93	0.90
Vocational degree	1	0.94
College degree	0.99	1.21

Notes: For the full model (without interaction effects) see Model 1a.

4.3 The Partner Effect Hypothesis

Our second hypothesis relates to the educational attainment of the **male partner.** Above, we argued that in a "male breadwinner context" like West Germany, one would expect that the partner's earning potential (measured by his educational attainment) fosters fertility. If one fails to control for the partner's characteristics, one might yield biased results on the role of women's education in fertility. This is of particular importance since more highly educated women often live with partners with comparable characteristics. In order to address this aspect, we add the partner's educational attainment to the model. Table 7 (Model 2a) shows that after adding the partner's characteristics, the impact of a woman's education completely vanishes, while the partner's educational attainment has a strong positive and significant effect. Having a partner with a college degree increases second birth risks by roughly 30 percent.

Table 7: *Event-history model of the transition to the second child, Model (2a): focus on the partner's education*

	β	exp(β)	t	
Baseline (ls)				
Intercept	-4.54		-33.71	***
0-3 years	1.22		27.85	***
3-5 years	-0.71		-19.98	***
5+ years	-0.12		-6.56	***
Age at first birth (ls)				
17-20	0.37		1.01	
21-25	0.06		3.00	***
26-30	0.02		1.98	**
31-35	-0.15		-9.30	***
Woman's education				
No degree	0.00	1.00	0.09	
Vocational degree	0	1		
College degree	0.03	1.03	0.50	
Partner's education				
No partner	-0.64	0.53	-11.86	***
No degree	-0.02	0.98	-0.34	
Vocational degree	0	1		
College degree	0.27	1.30	5.66	***
Log Likelihood		-7,840		

Notes: (1) Method: event-history model (generalized Gompertz)
(2) time variable: age of first child
(3) ***: $p \leq 0.01$ **: $0.01 \leq p \leq 0.05$ *: $0.05 \leq p \leq 0.10$
(4) ls= linear spline
Source: *Mikrozensus* 1997 (own estimates)

The woman and the partner's educational attainment combined

In a further step, we employ an interaction between the woman's and the partner's educational levels (see Table 9). The combination "both vocational degree" serves as the reference category. If it was solely the partner's education which mattered in the decision to have a second child, couples where the male partner has a college degree should have a particularly high second birth risk. If only the woman has a college degree, second birth risks should be rather low. The empirical analysis partially supports this hypothesis. If only the male has a college degree, second birth risks increase by roughly 30 percent (compared to the reference group of homogamous couples with vocational degrees). If only the woman has a college degree, second birth risks are slightly lower than for the reference category. However, "homogamous couples" with college degrees encounter a relative second birth risk of 34 percent (compared to the reference group). Obviously, in such cases it is not possible to separate the impact of the woman and her partner's educational attainment on the risks of having a second child. Here, it would be important to take into account other employment indicators (such as labor market income or labor market status). Unfortunately, data restrictions preclude us from investigating the role of the partner's employment situation any further.

Table 8: *Event-history model of the transition to the second child, Model (2b): combined effect of the woman and the partner's education*

	Woman's education		
	No degree	**Vocational degree**	**College degree**
	exp(β)	exp(β)	exp(β)
Partner's education			
No degree	0.99	0.95	0.96
Vocational degree	0.95	1	0.99
College degree	1.37	1.27	1.34

Notes: For the full model (without interaction effects) see Model 2a.

4.4 The Self-Selection Hypothesis

In a last step, we address the question of whether the positive effect of female educational attainment can be attributed to self-selection. Above, we argued that in an institutional framework where childrearing and employment is not compatible, women are basically confined to choose between a continuous employment career and motherhood. Women who choose to set up a family in such an institutional context should either have low employment ambitions or a high family orientation. This "self-selection" might bias the effect of college education on second birth risks.

In order to investigate this issue, we proceed in a manner suggested by Kravdal (2001). He estimates the transition to the first, second and third child within a joint model and adds a common unobserved heterogeneity factor. Similar to Kravdal (2001), we estimate first and second birth probabilities within a joint model and we insert a common **unobserved heterogeneity** factor to the model. In contrast to the study by Kravdal, however, we do not use event-history techniques for all birth parities. Instead, we simply estimate the probability of having a first child in a probit model; the transition to the second child is, however, estimated in an event-history model.

Below, we report the underlying functional relationship. The decision to have a first child is determined by

$$y = P(y) = \begin{cases} 1 \;\; if \quad first \quad child \\ 0 \;\; if \; childless \end{cases} \tag{2}$$

For the probit equation, Z represents the independent variables that influence the decision to have a first child, α_1 represents the intercept and α_2 the parameters. For the transition to the second child, $h(t)$ is the hazard of occurrence at time t, β_1 is the baseline hazard, X are the covariates and β_2 the respective parameters. The symbol σ represents the unobserved heterogeneity factor that is the same for both birth parities. σ is supposed to be normally distributed with mean 0 and a variance of σ^2.

$$\begin{aligned} y &= \alpha_1+\alpha_2 Z + \sigma \\ \ln h(t) &= \beta_1+\beta_2 X+ \sigma \end{aligned} \tag{3}$$

The model without unobserved heterogenity (Model 3a in Table 2) provides basically the same results as Model 2 (see above). It shows that the partner's educational attainment has a positive effect on second birth risks, while woman's college education plays no role in the transition to the second child.

Model (3b) contains the results after adding unobserved heterogeneity to the model. We basically yield similar results as Kravdal does for his sample of Norwegian women

Table 9: *Simultaneous estimation of probability of having a first child and the transition to the second child (Model 3a+b)*

	Model (3a)				Model (3b)			
	β	exp(β)	t		β	exp(β)	t	
Probit model (first birth)								
Intercept	0.89	2.44	39.40	***	1.68	5.35	7.02	***
Woman's education								
No degree	0.13	1.14	3.09	***	0.24	1.27	2.81	***
Vocational degree	0	1			0	1		
College degree	-0.54	0.59	-11.22	***	-1.00	0.37	-6.07	***
Partner's education								
No Partner	-1.06	0.35	-29.19	***	-1.98	0.14	-7.09	***
No degree	-0.17	0.84	-2.82	***	-0.32	0.73	-2.60	***
Vocational degree	0	1			0	1		
College degree	0.05	1.05	1.11		0.10	1.10	1.09	
Hazard model (second birth)								
Baseline (ls)								
Intercept	-4.54		-33.69	***	-6.68		-10.16	***
0-3 years	1.22		27.87	***	1.78		9.47	***
3-5 years	-0.71		-20.01	***	-0.40		-4.97	***
5+ years	-0.12		-6.55	***	-0.07		-3.08	***
Woman's education								
No degree	0.00	1.00	0.09		0.09	1.09	1.16	
Vocational degree	0	1			0	1		
College degree	0.03	1.03	0.53		-0.29	0.74	-2.21	**
Partner's education								
No partner	-0.64	0.53	-11.82	***	-1.74	0.18	-6.11	***
No degree	-0.02	0.98	-0.34		-0.14	0.87	-1.29	
Vocational degree	0	1			0	1		
College degree	0.27	1.31	5.68	***	0.47	1.60	4.73	***
Age at first birth (ls)								
17-20	0.37		1.02		0.28		0.59	
21-25	0.06		2.97	***	0.09		3.01	***
26-30	0.02		1.99	**	0.04		2.55	***
31-35	-0.15		-9.30	***	-0.20		-8.74	***
Sigma	**--**				**1.60*** **			
Log Likelihood			-12,399				-12,383	

Notes: (1) Method: simultaneous estimation of a probit model (estimating the probability to have a first child at the time of interview) and an event-history model (generalized Gompertz; time variable: age of first child).
(2) ***: $p \leq 0.01$ **: $0.01 \leq p \leq 0.05$ *: $0.05 \leq p \leq 0.10$
(3) ls= linear spline

Source: *Mikrozensus* 1997 (own estimates)

(2001). The unobserved heterogeneity component is positive and significantly different from zero. This means that there are unobserved respondent-specific characteristics which affect fertility decisions. Omitting this unobserved heterogenity component has important consequences. Most importantly, it biases the role of women's college education upward. The interpretation we suggest in this context is that "family-prone" college graduates select themselves into the sample of women at risk of first birth. This "family-proneness" also fosters their transition to the second child.

4.5 Summary of Empirical Results

Table 10 summarizes the results of the various models. It clearly shows how sensitive the effect of women's college education is to the specification of the model. It turns from positive and highly significant to small and negligible to significantly negative depending on the way the model is specified. Model (1a), which solely controls for the age at first birth, suggests that a woman's college education increases second birth rates by 14 percent (compared to women with only a vocational degree). After adding the partner's educational attainment, the effect of women's education vanishes; the partner's college education increases second birth risks by 30 percent. In the final "selection model", the effect of the partner's college education on second birth risks is 60 percent. Here the effect of a woman's college education becomes strongly negative. Compared to women with a vocational degree, second birth risks are reduced to three-quarters.

Table 10: *Summary of results*

	Woman's education				Partner's education			
	β	exp(β)	t		β	exp(β)	t	
Model (0)								
No degree	-0.06	0.94	-1.49		--			
Vocational degree	0	1			--			
College degree	0.10	1.10	1.80	*	--			
Model (1a)								
No degree	-0.06	0.95	-1.35		--			
Vocational degree	0	1			--			
College degree	0.13	1.14	2.33	**	--			
Model (2a)								
No degree	0.00	1.00	0.09		-0.02	0.98	-0.34	
Vocational degree	0	1			0	1		
College degree	0.03	1.03	0.50		0.27	1.30	5.66	***
Model (3a)								
No degree	0.09	1.09	1.16		-0.14	0.87	-1.29	
Vocational degree	0	1			0	1		
College degree	-0.29	0.74	-2.21	**	0.47	1.60	4.73	***

Notes: Model (0): No other controlling variables
Model (1a): The only other controlling variable is the age at first birth.
Model (2a): The other controlling variables are the age at first birth and the educational level of the partner.
Model (3a): Results of the simultaneous estimation of the probability to be childless and the transition to the second child including unobserved heterogeneity (for the full model, see above).
Source: *Mikrozensus* 1997 (own estimates)

5. Conclusions

In this paper, we analyzed the transition to the second child in West Germany using data from the 1997 German micro-census. In line with other studies, we found a positive effect of female education on second birth risks. We have argued that the positive effect of women's education on second birth risks is due to model misspecification. We investigated three aspects in this context. First, we investigated the hypothesis that more highly educated women are under a "time-squeeze" which accelerates the transition rate to the second child. Second, we argued that there is substantial educational homogamy in Germany. A positive effect of a woman's college education on second birth risk might be attributable to the omitted partner's educational

characteristics. Third, we hypothesized that more family-oriented college graduates are more likely to select themselves into the group of women at risk of second birth. In the empirical analysis, we estimated several event-history models for the transition to the second child. Our results can be summarized as follows.

- There is ambiguous evidence for the "time-squeeze hypothesis". As a first step, we applied models similar to those that were employed to investigate the transition to higher parity births in Sweden, Norway and Austria (B. Hoem 1996, Hoem et al. 2001, Kravdal 1992). Contrary to these studies which showed that the age at first birth confounds the effect of women's education on the transition to the second and third child, we did not find consistent support for this hypothesis for the West German case.
- The role of the partner's educational attainment turned out to play a significant role in the transition to the second child in West Germany. Furthermore, after controlling for the partner's educational attainment, the effect of the woman's education becomes insignificant. This finding complies with the expectation that in the institutional context of West Germany, the employment situation of the male "breadwinner" is crucial for the decision to have a larger family. Furthermore, the results also suggest that it is of vital importance to take into account the partner's characteristics. Failing to account for them upwardly biases the coefficients for female educational attainment.
- Apart from a strong confounding effect of the partner's characteristics, we also found support for the "selection hypothesis". After controlling for unobserved heterogeneity factors, the positive impact of women's college education becomes significantly negative. Our interpretation of this finding is that a "family-proneness" selects women into the sample who are at risk of second births. This self-selection biases the effect of college education on the transition to the second child.

Further analysis, which in particular takes into account the women's employment behavior after first birth and more detailed information on the partner's labor market situation, are certainly required for a more comprehensive analysis. Moreover, our interpretation of the unobserved heterogeneity component is based on crude speculations. The hypothesis of a "family proneness" is surely interchangeable with other hypotheses on unmeasured personal specific characteristics, such as differences in fecundity. Although we are unable to pinpoint the precise mechanisms working behind this "selection process", it is important to note that the model is **not** robust towards adding unobserved heterogeneity. One implication for future research is to seek richer data sets that would be able to tackle the factors which bias the results.

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Changes

On October 14th 2002, per request of the author, the following changes were made:

1. On page 18, the sentences "Table 1 exemplifies this issue by displaying the employment rates of West German mothers by the age of the youngest child (Note 2). When the youngest child is between 3 and 6 years old…"

were changed to:

"Table 1 exemplifies this issue by displaying the employment rates of West German mothers by the age of the first child (Note 2). When the first child is between 3 and 6 years old…

2. On page 19, the table head "Age of youngest child"

was changed to

"Age of first child."

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Appendix

Testing the Proportionality Assumption

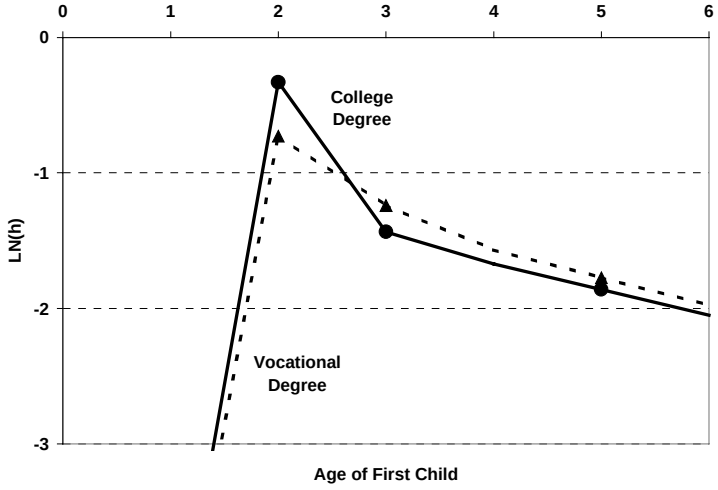
The survival curves (see Figure 2) show that college graduates have a second child more quickly compared to their non-college educated peers, but that they nevertheless encounter a fairly similar progression ratio to the second child. In order to address this issue in the event history model, we employ an interaction of the woman's educational level and the baseline hazard, i.e. the age of the first child. Table A1 displays the results and Figure 4 plots the estimated baseline hazards for women with a vocational degree and for women with a college degree respectively. The figure reveals that the hazard rates for college graduates are higher when the first child is below age three. When the first child is older than age three, though, the hazard rate for women with a vocational training certificate is the same. Although such a model more accurately models the baseline hazard, it is rather complex and makes it cumbersome to add additional interaction effects, which is why we did not use this model for the subsequent analysis.

***Table A1**: Event-history model of the transition to the second child, Model: interaction between the baseline and education*

	Woman's education		
	No degree	**Vocational degree**	**College degree**
	β	β	β
Baseline (ls)			
Intercept	-9.31	-9.23	-9.24
0-2 years	4.36	4.36	4.46
2-3 years	-1.16	-0.51	-1.11
3-5 years	0.08	-0.33	-0.24
5+ years	-0.29	-0.21	-0.19

Notes: (1) For the full model (without interaction effects) see Model 0.
(2) We added an additional node to allow the baseline to vary more flexibly.
(3) ls= linear spline

Figure A1: *Baseline hazard by educational attainment*



Notes: For the estimates, see Table A1.
Source: *Mikrozensus* 1997 (own estimates)

Notes

1. We restrict our analysis to the western states of Germany, i.e. the territories of the former Federal Republic of Germany (hereafter referred to as "West Germany"). We do not take the eastern states of Germany ("Neue Länder") into account. Although East Germans were subject to a similar tax and transfer system in the 1990s, there were still important differences in the institutional constraints and this would have required a separate analysis (see e.g., Kreyenfeld 2001).

2. We distinguish between not employed, part-time employed (defined as working less than 35 hours) and full-time employed (working more than 35 hours). We follow here a definition given by the Statistisches Bundesamt (1999).

3. There are further features, which we are unable to address in our analysis. For example, there might be differences in the use of contraceptive methods by educational level. On the one hand, it seems plausible that the use of more efficient contraceptive techniques should have a depressing effect on fertility. Ní Bhrolcháin (1988) shows, however, that better control over fertility can also have an accelerating effect on the spacing of children.

4. It is not possible to distinguish biological children from stepchildren (see also Kreyenfeld and Huinink 2001).

5. We are using a "whole-year time scale", i.e. the spacing of the children is measured by the difference between the year of first and of second birth. Related to this, we censor the cases at the end of 1996. The interviews were conducted in April, 1997, and we therefore do not have a full last year at our disposal.

6. One could argue that women are only at risk of a second birth if they have a partner. Therefore, one should censor the cases when respondents experienced the breakup of a partnership. However, the partner's characteristics only relate to the date of the interview and we are unable to locate the date of union disruption. We control for the presence of a partner at the time of interview, however, we do not interpret this variable in any substantial manner.

7. One could argue that the survival curves suggest that more highly educated women opt more rapidly for the second child, but that they nevertheless encounter a fairly similar final progression ratio to the second child. To fully take this aspect into account in the event history model, one would have to employ an interaction of the baseline with the woman's educational attainment (see Appendix).

8. Note that the coefficient for college educated women who had their first child at the ages 31-35 is negative. However, this is compared to the reference category, i.e. women with a vocational degree who had their first child at the ages 21-25. Compared to women with a vocational degree who had their first child at the ages 31-35 the coefficient is positive, first birth risks for college graduates increases here by 40 percent (0.89/0.64=1.39).
9. Instead of the mean age at first birth, we also used age quartiles but found no substantial change in the results.